# Firms' rents, workers' bargaining power and the union wage premium in France

Thomas Breda

PARIS SCHOOL OF ECONOMICS
ECOLE D'ECONOMIE DE PARIS

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PARIS-JOURDAN SCIENCES ECONOMIQUES
48, BD JOURDAN – E.N.S. – 75014 PARIS
TÉL. : 33(0) 1 43 13 63 00 – FAX : 33 (0) 1 43 13 63 10
www.pse.ens.fr

# Firms' rents, workers' bargaining power and the union wage premium in France

Thomas Breda[1]

Paris School of Economics

**Abstract:**

In this paper, I study the wage premium associated with firm-level union recognition in France and show that this premium is due to a rent-extraction phenomenon. Using a large matched employer-employee dataset from a 2002 survey in France, I first estimate a series of wage determination models that control for individual and firm-level characteristics. I find that union recognition is associated with a 2-3% wage premium. To show that this premium results from a non-competitive phenomenon, I construct a bargaining model and estimate it empirically using a smaller but very detailed matched employer-employee dataset for 2004. The model predicts in particular that the wage premium obtained by unions should increase both with their bargaining power and with the amount of quasi-rents available in the firms they organize. These predictions are validated empirically when I use the firms' market share as a proxy for their quasi-rents and the percentage of unionized as a proxy for the unions' bargaining power. All the results remain valid when I control for the firm-level workers' average productivity.

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## Introduction

Why are unionized workers paid more than their non-unionized counterparts? An obvious explanation, often called the "causal effect" of unions, is that unions raise wages through bargaining and rent extraction. But a wide range of alternative explanations are possible: unionized workers can be more productive than non-unionized ones (selection of unionized workers), organized firms can have unobserved characteristics correlated with higher wages (selection of organized firms) and wage gains for unionized workers can be counterbalanced by losses on other aspects (compensating wage differentials). Due to econometric limitations, studies are often unable to disentangle completely those various explanations. Typically, microeconomic studies based on a sample of workers may potentially confound bargaining status with other firm-level characteristics such as firm size. This is the case for a huge body of studies in the United States that finds sizeable union wage premiums[2]. However, more recently DiNardo and Lee (2004) used a regression discontinuity design technique to identify the "causal effect" of unions. Using a sample of U.S. establishments that changed union status as a result of a union certification election, they found no effect of union coverage on wages.

Consistent with the rent-extraction interpretation is the idea that the wage differential between unionized and non-unionized firms[3] should be increasing both with the amount of rent per worker available to the unionized firms and with the bargaining power of unions in these firms. In this paper, I derive these two predictions from a simple bargaining model and test it using a detailed linked employer/employee dataset from the French private sector. First, the data contains subjective information on the surveyed firms' market share. Under the assumption that firms declaring a high market share should have on average more rents per worker that unions can potentially extract than those declaring a low market share, I split the sample of firms in two groups according to their declared market shares. I then compare the *ceteris paribus* wage differential between unionized and non unionized firms obtained in these two groups. I argue that a higher differential observed in the group of high-market-share firms would strongly reinforce the rent extraction interpretation of the wage differential between organized and non organized firms. Second, France is a country of "open-shop"

[2] Studies that use a panel of workers from the Current Population Survey (CPS) cannot take in account firm's characteristics. See Lewis (1986) for an extensive survey of the early literature and Freeman and Medoff (1984) or Card (1996) for famous examples based on the CPS.

[3] In this paper I focus on the usual "union recognition wage premium", that is the wage differential between workers who are covered by unions at the firm level and those who are not covered.

unionism, with no requirement for workers to be unionized when a union is recognized in their firm. I argue that a larger proportion of unionized workers in a firm where a union is recognized indicates a higher support toward the union and thus a higher bargaining power of the union. The rent-extraction view of union wage differentials then predicts that the wage premium obtained by unions should increase with the proportion of unionized workers in organized firms. I take advantage of the information available on the proportion of unionized workers in the dataset I use to test empirically this second prediction. The workers' bargaining power is likely to be endogenous to the rents available in their firm (the higher their potential gains, the higher the incentive for workers to pay the cost to organize and bargain collectively). I thus estimate a structural wage equation derived from a simple bargaining model that models simultaneously the rents per worker available at the firm level and the workers' bargaining power. Finally, the workers' productivity is also likely to be endogenous to the rents available in their firm (more productive workers are more likely to generate higher profits and rents). To control for this possible selection effect, I use the workers average productivity at the firm level as an additional explanatory variable in some of my regression models.

A second important feature of this study is that it focuses on France, a country which has the reputation to have extremely powerful unions. According to an article by Craig Smith published in the New York Times in 2006[4], "Despite one of the lowest rates of unionization — only about 8 percent of the French work force are members — the unions have enormous leverage over the government. They play a unique organizational role in France's hierarchical society, rallying the populace accustomed to a confrontational relationship with leaders considered elitist. Spark-plug unions, some people call them." This commonly accepted view on the strength of French unions relies on evidence at the national level and on large national strikes or demonstrations occurring from time to time and largely advertized in the general media. But what is the strength of French unions at the firm level?

I answer this question by comparing the wage differential between organized and non organized firms I obtain in the private sector for France with measures of this differential obtained aboard. In particular, using a dataset similar to theirs, I reproduce the main empirical specifications of Card and De La Rica (2006) who studied the wage premium associated with firm-level contracting in Spain. As France and Spain are neighbor countries with similar

[4] See the following webpage for the entire article:
http://www.nytimes.com/2006/03/29/international/europe/29unions.html

industrial relation systems, the comparison should shed some light on the real strength of French unions at a decentralized level.

## Institutional Settings

The legal settings of union representation in France have been slightly modified on the 4th of May 2004 and more recently on the 20th of August 2008. As this study focuses on years 2002 and 2004, I describe the functioning of industrial relations before these two laws were passed. I begin with a brief description of industry-level bargaining and then turn to a more precise description of firm-level industrial relations.

At first sight, France shares with most continental Europe countries characteristics of a regulated industrial relation system with multi-level bargaining. First, industry wide agreements negotiated by unions and employer associations cover most of the workforce. Second, individual employers can sign firm specific agreements with unions when unions are recognized at the firm level. According to the Statistics Department of the French Ministry of Labor (DARES), 97.7% of the workforce was covered by a collective agreement in 2004. With a union density around 8%, France is the OECD country with both the highest coverage rate and the lowest union density (OECD Employment Outlook, 2004).

Industry-level bargaining is organized by branches. A branch is a bargaining unit regrouping workers in a same industry or group of industries, sometimes in a delimited region and sometimes also with a specific occupation[5]. When an agreement is signed in a branch between unions and an employer association, only the firms whose the employer is a member of the association are initially covered. An extension of the agreement for all workers in the branch can be asked by unions, the government or another employer association. The extension is made as soon as the agreement is proved conform to the general law[6]. In practice, the extension mechanism is very common (Barrat and Daniel 2002), which explains that most of the workforce is covered by industry-wide agreements.

In 1982, the Loi Auroux (August 4, 1982) encouraged decentralized bargaining. As a consequence, industry-level bargaining became less significant (Barrat et al 1996). Some of the existing agreements are even outdated because they have been rarely renegotiated in the

[5] For example, white collars workers in the construction sector bargain at the national level whereas other occupations bargain at the regional level (see Ayouvi-Dovi et al, 2009).

[6] This differs, on the one hand from Spain where industry-level agreements are automatically extended to the entire industry and, on the other hand from Germany where conditions of representativeness are also necessary for the extension (see Du Caju et al. 2008 for details).

past two decades and are now weaker than national standards in many sectors and regarding many topics. In 2006, exactly 50% of the 160 branches covering more than 5,000 employees[7] had a branch minimum wage which was below the national minimum wage and was consequently useless. Figure 1 illustrates this point and plots the French national minimum wage in 2006 as well as the distribution of the 160 largest branch minimum wages. To summarize, almost all workers are covered by industry-level agreements (which render impossible, in the absence of a comparison group, the identification of the effect of these contracts on wages) but a lot of these contracts are weak or even outdated, which should leave room for unions to bargain at the firm level.

Making comparisons between the degree of bargaining at the national, industry and firm level, the 2004 OECD Employment Outlook classifies France in the second group of OECD countries with the most decentralized bargaining institutions (with Australia, Italy, the Slovak Republic and Switzerland) just behind the U.S., the U.K., Canada, Poland, Korea and Japan. Regarding this classification, I only focus on the union wage premium at the firm level, similarly to the approach taken in Anglo-Saxon studies (where bargaining in the private sector is only decentralized) and by Card and De La Rica for Spain (2006). In that sense, my approach differs from a recent literature on continental Europe countries which focuses on industry-level bargaining or examines the relative influence of the different levels of bargaining on the overall structure of wages (Ayouvi-Dovi et al 2009, Cardoso and Portugal 2005, Fitzenberg et al 2007, Plasman et al 2007, Rusinek and Rycx 2008).

Firm-level agreements can be signed between unions and employers as soon as unions have been recognized within firms. Concerning wages, these agreements can only improve the industry minimum wage and must be above the national minimum wage. Three key features differentiate France from Anglo-Saxon countries concerning union coverage at the firm level: first, there is no certification election, second, many unions can be present in the same firm and represent workers collectively and third, unionism is completely “open shop”.

*There is no certification election:*

To be recognized in a firm with more than 50 employees, the main unions almost only need to find one worker who accepts to officially represent the union in the firm. Such a

[7] There are about 700 branches in total. The Ministry of Labor provides information on the 160 that cover more than 5,000 employees each. In total, these branches cover more than one half of the private sector total employment.

worker is called a union representative. Table 1 presents a brief description of the main French unions and gives the distribution of the union representatives in terms of the unions they belong to. We can see that more than 95% of union representatives belong to only 5 large national "historical" unions. These "historical" unions are recognized as legal bargaining units within firms as soon as a worker accepts to be their representative[8]. This is a fundamental feature of the French industrial relations: there is no certification election required for historical unions to organize larger firms. In firms with size between 10 and 50 employees, unions have to choose their representatives among workers who have already been elected, the so-called "firm delegates". These "firm delegates" are legally recognized non-union representatives acting as the voice of the workers in their day-to-day relationship with the employer (they are generally also members of the work councils). They are elected every four years by workers in firms with more than 10 employees among voluntary candidates in a simple majority rule voting (the winning candidates are simply those that have collected the larger number of votes). The process of union recognition is more binding in firms with size between 10 and 50 employees, but even in these firms, union recognition remains less binding than the U.S. certification process which requires a majority of workers to be pro-union. The very weak legal constraints bearing on firm-level union recognition makes it easier for unions to legally organize firms and get a legal framework to bargain over wages officially. However, the low organizational cost paid by unions in these firms and the fact that they are not necessarily supported by the majority of the workforce should limit their bargaining power and the scope of their action.

*Different unions can organize the same firm:*

The recognition process described above applies to each union, which makes in theory unlimited the total number of unions that can cover the workers of a given firm. Table 2 shows the distribution of the workplaces in terms of the number of unions present in them. The second column gives the non weighted distribution in the dataset I use- the REPONSE data described in the next section- whereas the third and fourth columns are obtained using weights that make the data representative of French private sector workplaces with more than 20 employees or of the workers in those workplaces. It can be derived from table 2 that around 36% of the private sector workplaces with more than 20 employees are organized, which represents 64% of the workforce in these workplaces. This discrepancy is explained by

[8] The other non historical unions might have to win a certification election to be recognized at the firm level if the employer or a worker asks for it.

the fact that the firm's probability to be organized increases considerably with its size (see table A1 in appendix A).

*Unionism is completely "open shop":*

When one or more unions are recognized in a firm, in place and newly hired employees do not have the duty to become union members, neither to participate in strikes. This enables me to use the percentage of unionized workers at the establishment-level as a measure of the unions' bargaining power. Finally, union contracts must apply to all workers in the firm. For this reason, I will study the effect of unions on both the wages of unionized and non-unionized workers.

Finally, the institutional settings concerning industrial relations and bargaining at the establishment level are exactly identical to the institutional settings at the firm level which is described above. As it appears to be more relevant, I conduct the empirical analysis of the effect of union recognition on wages at the establishment level[9].

## Data description

The empirical analysis relies on two sources of data. First, the 2002 French Wage Structure Survey (ESS02) collected detailed salary and job information for up to 60 employees in each of some 15,000 private sector establishments in the manufacturing, construction, trade and service industries. The design of the survey allows to model wage outcomes at the employee level while including controls for establishment characteristics. Agriculture, mining, and household services are missing from the ESS02 sample as are small establishments (less than 10 employees). As firm-level union coverage is extremely low for small workplaces and in the industries missing from ESS02, the limited coverage of the ESS02 is not a major problem for my study. I have excluded from the sample chief executives as well as workers having their wage in the first and last percentiles of the hourly wage distribution.

The second dataset I use is the 2004 French Workplace Employment Relations Survey (REPONSE04) conducted by the Ministry of Labor towards up to 10 employees in each of 2929 business establishments with more than 20 employees. REPONSE04 contains extensive

[9] It is difficult to know exactly what the actual bargaining unit is. For mono-establishment firms, establishment-level and firm-level union recognition are of course confounded. Multi-establishments firms are large enough to always have in practice unions recognized at the firm-level. For these firms, only establishment-level union recognition varies enough to offer a matter of comparison.

information on industrial relations at the workplace level and on the firms' organizational and technological structure. In each surveyed workplace, union density, the name of the unions that are present and the existence of a firm-level contract are available. I will use union density to proxy the union's bargaining power. REPONSE04 also contains information on the market share of each establishment, as declared by its manager. I will use this information to proxy the firm's market power and potential rents. Net hourly wages in December 2003 have been retrieved from Social Security records (the Déclaration Annuelles de Données Sociales, DADS) by the Ministry of Labor for the workers surveyed in REPONSE04 and have been matched with the dataset. I have also excluded from the data sample chief executives[10]. The REPONSE04 survey covers mainly the private sectors but some public companies are also present, as well as non-profit associations and cooperative firms. Since this paper focuses on unions and rent-sharing, I have removed these observations and kept a final sample of 2451 business establishments owned by private non cooperative firms.

Comparing to ESS2002, the main inconvenience of REPONSE04 is that it is relatively small, and its main advantage is to contain extensive workplace-level information. I use ESS2002 to estimate precisely the cross-sectional union wage gap and make comparison with similar studies and REPONSE04 to test the more elaborate predictions that these union wage gaps should increase with firms' market shares and workers' bargaining power if they are due to rent extraction[11].

## The union wage premium in a standard wage determination model

Before turning to a more sophisticated econometric analysis that aims at capturing the causal effect of unions on wages, I provide a precise estimation of the union wage premium that controls for individual-level and establishment-level observable characteristics. To do so, I present a series of regression models of the type:

$$\log(w_{ij}) = X_i\beta + \mathrm{Z}_{\mathrm{j}}\gamma + U_j\alpha + \varepsilon_{ij} \tag{1}$$

[10] Since wages come from an administrative source, I have not excluded workers having extreme wages. However, I have also performed the whole empirical analysis both on the full and truncated samples (removal of 0.5% or 1% tails of the wage distribution) of both the ESS02 and REPONSE04 datasets. The results (available on demand) are always very close.

[11] The two datasets I use have twins in other countries that have been used a lot to study unions. REPONSE follows the same design than WERS in the U.K. See Bryson et al. 2009 for a study that uses both REPONSE and WERS to study unions and workplace. Wage Structure Surveys similar to ESS have been used by Plasman et al (2006) to study the effect of multi-level bargaining on wages in Belgium, Denmark and Spain and Card and De La Rica (2006) in Spain.

where $w_{ij}$ represents the hourly wage of individual i in establishment j, $X_i$ is a set of observed skill characteristics (such as age and education) of worker i, $Z_j$ a vector of firm-level covariates and $U_j$ an indicator for the presence of one or more unions in establishment j. Assuming that $E[\varepsilon_{ij}|X_i, Z_j, U_j]=0$, the effect of establishment-level union recognition can be estimated consistently by a conventional (OLS) regression applied to (1).

The first 3 columns of table 3 present a series of regression models following equation (1) on the ESS02 dataset. In the first column (specification 1), only a dummy for union recognition at the workplace level is included. The estimated coefficient is just over 20% suggesting a large premium associated with union recognition. As shown by the results in column 2, more than 80% of this gap is explained by differences in the characteristics of workers and firms between unionized and non unionized workplaces. The covariates in this specification include the individual worker's age, education and occupation (both divided in 4 groups), a dummy indicating whether he or she was employed on a temporary contract, and dummies for establishment size, occupation, industry, and region. Many of the control variables are highly statistically significant, and their inclusion raises the R-squared above 60%.

Estimating a standard wage determination model, Card and De La Rica (2006) found a wage premium of around 12% for women and 8% for men for firm-level contracting. To describe their results, the authors explained that "these models are very similar to the specifications fit in many previous studies of wage determination in the United States, the United Kingdom, and continental Europe, and yield estimated premiums for firm-level bargaining that are comparable to (or little smaller than) the unionized wage premiums typically estimated in the United-States". Specification 2 of table 3 presents the results of a similar wage determination model for France but shows that the wage premium associated with establishment-level union recognition is only around 2.5%[12]. This premium is a lot lower than what is found with this kind of cross-sectional approach in most developed countries[13].

[12] Specification (2) of table 3 tries to reproduce very closely specification (2) of the third table in Card and De La Rica. In their specification, they included 2 additional controls for the market orientation and public ownership status of the firms that are not available in ESS02 and produced estimations for men and women separately. They control for years of education and I use 4 education dummies, they have 16 indicators for regions and 6 for industries, I use 10 indicators for regions and 9 for industries. The other control variables are rigorously identical. When I produce separate estimations for men and women as they do, I obtain a slightly higher coefficient for women than for men (the regression coefficient is 0.028 for women and 0.026 for men). For detailed studies of the effect of union recognition on men and women in France, see Leclair and Petit (2004) and Duguet and Petit (2009).

[13] The recent study by Blanchflower and Bryson (2010) finds a union wage premium around 5% in UK (private sector, years 2001-2006). This is lower than previous estimations for UK but still at least twice larger than my estimates for France.

Finally, column 3 of table 3 presents the results of a regression model with an extended set of control variables: 2-digit industry dummies as well as 10 dummies for age and 4 dummies for tenure have also been included as controls. The wage premium associated with union recognition at the workplace level is reduced by one additional third (comparing with column 2). Overall, the results of table 3 are close to those found in older firm-level data studies in France (Coutrot, 1996; Laroche 2004). As a robustness check, I have reproduced in appendix table A3 the wage models estimated with ESS02 in table 3 using the alternative dataset REPONSE04 and I find very similar results.

In their study of Spain, Card and De La Rica (2006) looked at the effect of firm-level contracting (rather than union recognition) on wages. Results for firm-level contracting[14] in France are presented in columns 4 to 6 of table 3. The estimated coefficient for firm-level contracting is not statistically significant and very close to 0 in specifications that include controls for workers and firms characteristics. The reason why I use union recognition rather than firm-level contracting is that reaching an agreement at the workplace or firm level is not necessarily the sign of strong unions. Unions can obtain a wage rise (through the threat of strike for example) without signing any agreement. When unions' demands are very strong for instance, the employer might only concede part of it, still leading to wage rises but with no final agreement reached. Also, multi-unionism put unions in competition at the firm level. In 2004, an agreement was considered legally valid as long as it was signed by one union in the firm. As a consequence, when more than one union is recognized in a firm, the employer might try to reach an agreement with the weakest union, leading to actually smaller wage rises than when the agreement is not reached. A practical example of this situation is illustrated by the CGT union (see table 1). This union is the most combative of the large French unions. By tradition, it signs very few agreements. Nevertheless, Breda (2008) showed with firm-level data that CGT obtains the largest wage rises.

Why is the union wage premium for France, while this country is supposed to have powerful unions, so much smaller compared to that in most other developed countries? A first explanation is the existence of a high and binding national minimum wage in France. In 2007, 12.9% of French workers were paid the national minimum wage (Berry, 2008). This high national minimum wage may simply leave little room for further bargaining at a decentralized level. A slightly different explanation relies on the work by Aghion et al (2008). France has evolved towards an equilibrium (in terms of industrial relations) with a highly regulated

[14] The variable "Firm wage contract" is an indicator of the signing of a new workplace or firm-level contract on wages in 2002. It is equal to 1 for 62.6% of the observations in ESS2002.

minimum wage and poor labor relations. In their view, the state regulation of the minimum wage crowds out the possibility for workers and employers to experience negotiation and develop trustful labor relations. If we suppose that wage rises at a decentralized level are more likely to be obtained when labor relations are good, then a high degree of regulation of the minimum wage is a substitute for good labor relations and thus for high wage rises at the decentralized level.

A second explanation for the low union wage premium in France directly derives from the analysis of the French institutional settings. In France, the large national unions are de facto recognized in firms or workplaces as soon as they find a worker who accepts to be their representative. This is a very weak legal constraint which implies in particular that a union can be legally recognized in a firm even though a large majority of the firm's workers are in fact against the union. In this case, the union cannot credibly threaten to begin a strike and its bargaining power will certainly be lower, leading to a lower wage premium. Since the cost to do so is low, unions have also an incentive to organize a large number of firms rather than just selecting those with a very high amount of rents. Table 2 indeed shows that, despite a low unionization rate, unions are present in a large number of firms. The small average premium associated with union recognition at the decentralized level has thus to be put in the context of the relatively large number of workers who benefit from such a premium. Finally, if the average quantity of rents available in unionized firms is lower than in other countries because unions have selected a larger number of firms, the average wage premium unions can extract is also lower. However, if this explanation holds, the union wage premium should be higher - maybe comparable to what is found in most developed countries- in firms with high potential rents.

I now push the analysis one step further and build a bargaining model to show that the union wage premium associated with union recognition is due to a rent-extraction phenomenon and is indeed higher in firms with high potential rents.

## Construction of a bargaining model

The larger the rents and the workers' bargaining power in a given firm, the higher their wages. In this section, I formalize this assumption in a simple bargaining model and derive wage equations to be estimated empirically. The goal of this more structural approach is to give evidence that the union wage premium is indeed due to bargaining and rent extraction, rather than selection effects or compensating wage differentials. This is done by deriving two simple testable predictions compatible with the rent extraction story, but much harder to explain if one believes that only selection effects and compensating wage differentials are at play in the union wage premium.

I first assume that in the absence of unions in her firm, worker i in firm j is paid a market hourly wage $w_{ij}^{m}$ that depends on her characteristics and on her firm's characteristics. Keeping the notation of the previous section, we have, for workers in non-unionized firms:

$$\log(w_{ij}^{m}) = X_i\beta + \mathrm{Z}_{\mathrm{j}}\gamma + \varepsilon_{ij} \qquad (2)$$

A prominent literature (Abowd and Lemieux 1993; Abowd and Allain 1995; Blanchflower, Oswald and Sanfey 1996) has shown that a lot of rent-sharing occurs in U.S., Canada and France. I nevertheless suppose in equation (2) that rent-sharing does not happen at the establishment-level in the absence of unions. Regarding the French law, actual firm-level bargaining (face to face discussion between the employer and a worker representative) can indeed only happen when unions are recognized. But one could think that implicit bargaining could still occur in non-unionized firm, leading to some rent-sharing. The existence of industry-level bargaining in France, even if weak (see above), might also imply some rent-sharing in non-unionized firms. Kramarz (2008) estimates a bargaining model with a large longitudinal dataset for France and shows that there is no rent-sharing in firms in which official bargaining does not take place[15], that is in firms in which unions are not recognized. To control for potential rent-sharing at the industry-level, I will nevertheless include in the firm's covariates $\mathrm{Z}_\mathrm{j}$ detailed industry indicators. I will also provide empirical evidence consistent with the fact that there is no rent-sharing in non-unionized firms in the next section of this paper.

When unions are present in a firm, each worker's wage $w_{ij}^{U}$ is the result of a Nash bargaining between the employer and the workers. Each worker's outside option in the bargaining is the market wage she could get in a non-union firm. The firm threat point is zero profit. Let us denote by $w_j^m = \sum_{i\in j} w_{ij}^m$ the threat point of firm j workers taken as a whole and $w_j^U = \sum_{i\in j} w_{ij}^U$ the total wage bill in firm j. The bargaining consists in maximizing the product of the employer and the workers surplus respective to their threat points:

[15] More precisely, he shows that 50% of quasi-rents are captured by workers in firms with official bargaining on wages and employment, whereas in firms with no official bargaining or official bargaining on wages only, there is no rent-sharing.

$$w_j^U = Arg\max(w_j^U - w_j^m)^{\varphi}(pF(L_j) - w_j^U)^{1-\varphi} \tag{3}$$

where $L_j$ is firm j labor force and $F(L_j)$ is its production function, while $p$ is a revenue shifter. $pF(L_j) - w_j^U$ are firm j profits. φ is the union bargaining power. The goal of this paper is not to make a detailed analysis of the various bargaining models, since it has already been done extensively in the literature (Abowd and Lemieux 1993; Blanchflower, Oswald and Sanfey 1996; Kramarz 2008). Yet some clarification is necessary. In the strongly efficient bargaining model (Brown and Ashenfelter 1986), the union and the firm bargain both on wages and on employment. In the weakly efficient bargaining model[16], the firm and the union bargain over wages only, while the firm unilaterally sets employment to its profit-maximizing level given the negotiated wage rate. Since it does not set out the arguments of the maximization, equation (3) is compatible with these 2 models. Abowd and Lemieux (1993) show that in the 2 models cited above, the solution of equation (3) is

$$w_j^U = w_j^m + \phi_j QR_j L_j \tag{4}$$

where $\phi_j$ is equal to $\varphi_j$ in the strongly efficient bargaining model and to a positive fraction of $\varphi_j$ in the weakly efficient bargaining model. $QR_j = (pF(L) - w_j^m)/L$ are the quasi-rents per worker in firm j and represent the profit per worker the firm would make if all the workers were paid their market wage. Equation (4) gives the share of quasi-rents going to the workforce. To know what each worker gets individually, it is necessary to make an assumption on how the union splits the bargained surplus between the firm workers. I make the usual assumption that the union is egalitarian and splits the surplus equally between all the workers. Under this assumption, equation (4) can be rewritten at the individual level:

$$w_{ij}^U = w_{ij}^m + \phi_j QR_j \tag{4'}$$

This simply means that the wage of worker i in firm j is equal to her individual market wage plus a share of the bargained surplus which is equal for all workers in firm j. Taking the log of equation (4'), we obtain $\log(w_{ij}^U) = \log(w_{ij}^m) + \log(1 + \phi_j QR_j / w_{ij}^m)$. Since firms' quasi-rents $QR_j$ are usually small relative to their total labor cost and since the workers bargaining power $\phi_j$ rarely exceed 0.5 (Kramarz 2008), we can work with first order terms:

$$\log(w_{ij}^U) = \log(w_{ij}^m) + \phi_j QR_j / w_{ij}^m \tag{5}$$

[16] This model is a version of the right-to-manage model or labor demand model (dating back to Dunlop, 1966) which includes bargaining on wages in the first step of the model rather than unilateral setting of the wage level by the union.

Substituting $w_{ij}^{m}$ by its expression in equation (2) and denoting by $U_j$ an indicator equal to 1 when unions are present in firm j, we finally get a general wage equation for both workers in union and non-union establishments:

$$\log(w_{ij}) = X_i\beta + \mathrm{Z}_j\gamma + U_j(\phi_j QR_j / w_{ij}^{m}) + \varepsilon_{ij} \qquad (6)$$

Equation (6) leads to 3 predictions for the wage premium $\phi_j QR_j / w_{ij}^{m}$ associated with firm-level union recognition:

*Prediction 1:* the greater the firm quasi-rents, the larger the union wage premium.

*Prediction 2:* the larger the union bargaining power, the larger the union wage premium.

*Prediction 3:* the larger a worker's (market) wage, the smaller his log-wage premium.

Predictions 1 and 2 directly come from the Nash bargaining framework. The idea is to use the large amount of information available in the REPONSE04 dataset to provide reasonable proxy variables for the firms' quasi-rents and union bargaining power and to empirically test these predictions. Prediction 3 however comes from an additional hypothesis on the objective function of unions: if unions are egalitarians, they split the bargained surplus equally between workers and thus, the larger a worker wage, the smaller the share of this fixed bargained wage premium in his total compensation. Prediction 3 leads to study the impact of unions on the intra-firm structure of wages. As this prediction is not a key feature of the bargaining model and can less arguably be used to prove that bargaining is indeed at play, I only test it in appendix B by running quantile regressions on the large sample dataset ESS02.

## Firms' rents, workers' bargaining power and the union wage premium

I first introduce the 2 proxy variables I uses for firms' quasi-rents and workers' bargaining power and test predictions 1 and 2 separately. A full estimation of equation (6) and a discussion of selection issues will follow in the next section.

According to prediction 1, if union wage premiums are due to bargaining, the larger a firm's quasi-rents, the larger these premiums. The *ex ante* quasi-rents on which the bargaining really occurs are not observable. What is observable in the data is the *ex post* result of the bargaining (accounting wages and profits). To recover a measure of quasi-rents, authors like Abowd and Lemieux (1993) or Kramarz (2008) have used an estimated market wage for each worker (rather than their actual wage) to compute the *ex ante* profits on which the bargaining occurs. Since this measure of quasi-rents remains highly endogenous[17], these authors also instrument it using measures of foreign competition shocks. An alternative strategy proposed by Blanchflower, Oswald and Sanfey (1996) is to use past profits at the industry level rather than current profits[18].

In this paper, I follow a more direct strategy and use a new and simple indicator of the existence of potential rents at the establishment level. This indicator is the establishment's market share as declared by its manager[19]. In the REPONSE04 survey, managers are asked if the market share of their establishment is lower than 3%, between 3% and 25%, between 25% and 50% or larger than 50%. Table 4 (first and second rows) shows the distribution of this subjective market share variable across the 1861 REPONSE04 establishments for which it is available. The establishments' market share is a direct measure of their market power, that is of their ability to raise unilaterally their sales price and profit margin. It is consequently a good measure of the *ex ante* potential rents firms can get in their industry, relative to their competitors. Of course, a firm's market share depends in the long run on its performance and might be correlated to the quality of its employees. But the market share varies little and is not affected in the short run by variations of wages, contrary to profits that are mechanically correlated to wages. Hence the use of the market share avoids some of the endogeneity problems emerging when using measures of quasi-rents derived from accounting data which in fact represent the *ex post* result of a potential bargaining.

Firms' market shares and labor costs are far less volatile and sensitive to economic shocks than profits or sales. The market share can be viewed as an indicator of the long-run firm health. We know from the theory of implicit contracts (Azariadis, 1975) that firms insure their workers against economic fluctuations (Guiso et al, 2005). This makes wages rigid in the short run and implies that the short-term relationship between current profit flows and current wages is a weakened measure of the total quantity of rent-sharing within firms. Indeed, if the bargaining occurs in the long run (as in a repeated game), the workers will want to exchange wage insurance in bad years against less rent sharing in good ones. In other words, the degree of rent-sharing in a given year might depend on the firm performance in the previous years. For this reason, studies that try to link directly profits to wages also have to deal with delicate

[17] Profits or alternative measures of quasi-rents derived from accounting variables such as sales can be endogenous in many respects. For example, in the efficiency wage theory (Akerlof and Yellen 1986), higher wages lead to higher profits rather than the contrary, leading to reverse causality in wage-profit regressions.

[18] There are also several papers that simply link current wages to current profits. See Fakhfakha and FitzRoyb (2002) for an example on French data and a review of the literature.

[19] The use of such a market share indicator is actually not completely new if we consider a broader strand of the literature: see Araï, Ballot and Skalli (1996) who used a similar variable in a different context.

framing problems (see Abowd and Lemieux 1993) and need to make assumptions on which profits are bargained in a given year (previous year profits, current profits, average past profits, etc). However, the use of the market share as an indicator of firms' potential rents captures the long-term firms' capacity to raise wages and avoids these delicate framing problems as well as biased relationships between wages and profits that can appear in the short run.

A common problem with objective measures of market share is that they require specifying the geographic units and industries to which firms belong. Indeed an objective measure of a firm market share is generally obtained by dividing its sales by the total sales in its industry and country. But some firms are not directly in competition with all other firms in their industry. More problematic, depending on their activity, firms operate at very different geographic scales. As proof of this, table 4 (last 2 rows) provides a distribution of the firms in the sample according to their declared targeted market. Only 24% of establishments operate on the national market. For them the standard market share indicators computed at the national level would really include the true competitors. For the remaining 76% of firms, these standard market-share indicators are inaccurate measures of the real competitive pressure that firms face. The subjective measure I use is not subject to these drawbacks since the interviewed managers should easily evaluate the real size of their market. Finally, my approach uses a measure of each establishment market share rather than the broader measure of industry concentration used in other studies (Blanchflower 1986). Since the degree of concentration of an industry is not informative of the relative market power of each particular firm in this industry, it seems inappropriate for the within-industry comparison of unionized and non-unionized firms I attempt in this paper.

The first 2 models of table 5 test the relationship between the market share and the union wage premium using the REPONSE04 data[20]. Model (1) uses the same workers and establishment control variables as Card and De la Rica (2006) in their study of Spain and model (2) adds more detailed controls for workers' age and tenure and establishment age, as well as detailed industry dummies. The establishment market share (grouped in 4 categories), union recognition and their interaction are the variables of interest in these models. A higher market share in the absence of unions is associated with lower wages in both models. A possible explanation for this is that firms with a high market share are also monopsonies in their industry. They employ specialized workers that have no or little outside options in other

[20] Descriptive statistics on REPONSE04 variables are available in appendix A and estimations of the union wage premium with various sets of control variables are presented in table A3 in appendix A.

industries and in the other firms of the same industry. These workers are paid less as a consequence of this when unions are not present to defend them and bargain over the rents available in those firms. Union recognition in low market share establishments is associated with lower wages in both models but the estimates are imprecise and are not statistically significant at conventional levels. This is consistent with the idea that in firms facing important competitive pressures (those with low market share), there are no rents for unions to bargain over. If any, wage increases obtained by unions in these firms should rise production costs above their competitive level and drive the firms out of the market, making them invisible in our data sample. Finally, the interaction between union recognition and market share is estimated to increase wages by about 2.5% in both models, in accordance with prediction 1. This effect is robust in model 2 to the inclusion of detailed industry fixed effects (161 dummy variables). The idea in model 2 is to identify the union effect based on intra-industry comparison of establishments with various market shares and union-recognition status. This is superior to model (1) since unions are historically better implanted in specific industries (such as manufacturing) and the average level of wages varies a lot across industries (Krueger and Summers, 1988) as well as the average degree of concentration. Nevertheless, since the REPONSE04 sample is relatively small, it makes sense to test the model's predictions using also the less demanding specification of regression model (1) which includes fewer covariates.

Models (1) and (2) in table 5 impose a linear increase of the union wage premium with the market share variable[21]. This assumption is relaxed in figure 2 which plots the union wage premium in each market share group, conditional on the detailed set of covariates included in model (2) of table 5. The union wage gap varies almost linearly between a non statistically significant -3% gap among the establishments declaring a market share smaller than 5% and a highly significant gap of 7.5% among the establishments declaring a market share larger than 50%. The difference between the union wage gaps in these 2 groups is thus 10.5%. Fischer's test of equality of the estimated union wage gaps in the first and last market share groups has a p-value of 0.003, implying that the union wage gap is almost surely larger among the firms with a large market share.

The proxy variable I use for unions' bargaining power is the percentage of unionized workers in the establishment. France is a country of "open-shop" unionism and workers do

[21] Remember that the market share variable is categorical. The fact that the union wage premium is linear with the market share variable does not mean that it is linear with the numerical value of the market share, which is actually not the case (see figure 2).

not have to be union members when a union is recognized in a firm even though union contracts legally cover all the workers. As a consequence, the percentage of unionized workers provides a direct indicator of the number of workers that supports the union(s) recognized in their establishment. Unions have more credibility to bargain and threat strikes if they are supported by a large number of workers. In this respect, the percentage of unionized workers is a good indicator of the unions' bargaining power. The percentage of unionized workers as declared by the establishments' managers has been bracketed in a 5-values variable. This variable is described in table 4[22].

The relationship between the unionization rate and the union wage premium is tested in models (3) and (4) of table 5. The set of control variables in these models is identical to those used in models (1) and (2). In both models (3) and (4), the interaction between union recognition and the unionization rate has a significant impact on the hourly wage, in accordance with prediction 2. Union recognition alone or having a high unionization rate without unions does not affect wages. This is an indication of the validity of the assumption made in the bargaining model that no bargaining occurs in firms where unions are not present. Indeed, if bargaining also takes place in non unionized workplace, we should probably observe a wage premium in non-unionized establishments having a lot of unionized workers. The right panel of figure 2 displays the estimated union wage gap in each unionization rate group. The union wage gap increases from virtually 0 in the group of establishments with less than 1% of unionized workers up to 12% among establishments having more than 10% of unionized workers. Fischer's test of equality of the estimated union wage gaps in the first and last unionization rate groups has a p-value of 0.012, implying that the union wage gap is larger among the firms with a large unionization rate at the conventional 5% statistical level.

## Estimation of the full bargaining model

I now turn to the structural estimation of a standard bargaining model with proxy variables for both quasi-rents and workers' bargaining power. Consistent with the assumption

[22] Note that only 34% of establishments where unions are recognized have more than 10% of unionized workers. The average unionization rate in establishments where unions are recognized is in fact very low (simple calculation from national statistics shows that it is less than 20%) and far lower than in US where 92% of the covered workers are unionized (Eren, 2009).

that there is no rent-sharing in workplaces where unions are not present[23], I define workers' bargaining power $\psi_j$ by $\psi_j = U_j \varphi_j$. Equation (6) can then be rewritten:

$$\log(w_{ij}) = X_i\beta + \mathrm{Z}_j\gamma + \psi_j QR_j \big/ w_{ij}^m + \varepsilon_{ij} \qquad (6')$$

To avoid the presence of the individual market wage $w_{ij}^m$ that comes in the right hand-side of the log-wage regression when we suppose that the union is egalitarian, a wage equation similar to equation (6') can also be estimated[24]:

$$w_{ij} = X_i\beta' + \mathrm{Z}_j\gamma' + \psi_j QR_j + \varepsilon_{ij}^{'} \qquad (6'')$$

I estimate equations (6') and (6'') derived from the bargaining model using two sets of proxy variables for quasi-rents and bargaining power. My first approach is to summarize the four-category market share variable in a "High Market Share" dummy variable (HMS) equal to 1 for establishments with a market share higher than 50% and 0 otherwise. I then postulate a linear relationship between quasi-rents and this high market share variable:

$$QR_j = aHMS_j + b + \mu_j \qquad (8)$$

with $a > 0$ and $\mathrm{E}[\mu_j | HMS_j] = 0$. The advantage of using a dummy variable rather than the four-category market share variable is that the linear relationship above can be supposed without loss of generality. The inconvenience is that the way to aggregate the market share categories is somehow arbitrary. I have chosen to isolate the establishments with a market share higher than 50% because the union wage gaps in figure 2 are clearly larger among this group.

Similarly, in figure 2, the union wage premium jumps up among the groups of establishments with a unionization rate above 10%. I thus approximate the workers' bargaining power by an indicator of High Bargaining Power (HBP) equal to 1 for unionized establishments having more than 10% of unionized workers:

$$\psi_j = cHBP_j + d + \eta_j \qquad (9)$$

with $c > 0$ and $\mathrm{E}[\eta_j | HBP_j] = 0$. Using equations (8) and (9), the wage equation (6'') can be rewritten:

$$w_{ij} = X_i\beta + \mathrm{Z}_j\gamma + \alpha_1 HMS_j + \alpha_2 HBP_j + \alpha_3 HMS_j HBP_j + \alpha_4 + \varepsilon_{ij}^{'} \qquad (10)$$

[23] This result has been proven by Kramarz (2010). The fact that the unionization rate alone does not lead to higher wages (table 5, columns 3 and 4) also confirms it. Finally, this assumption is also consistent with the French legal settings that make actual bargaining legal only when unions are recognized.

[24] Equation (6'') is derived from equations (4') considering that $w_{ij}^m = X_i\beta' + \mathrm{Z}_j\gamma' + \varepsilon_{ij}^{'}$.

with $\alpha_1 = ad, \alpha_2 = bc, \alpha_3 = ac$ and $\alpha_4 = bd$. Equation 10 is estimated in models (1) and (2) of table 6 using 2 different sets of covariates. In the most precise specification (model 2), the estimate of $\alpha_3$ is statistically significant at the 5% level and equal to 1.78. $\alpha_1$ is estimated to be significantly negative whereas $\alpha_2$ is virtually equal to 0. The negative estimate of $\alpha_1$ would imply that $d$ is negative, which would in turn lead to suppose from equation (9) that the true workers' bargaining power $\psi_j$ is on average negative in establishments where $HBP_j$ is equal to 0. Since this bargaining power parameter is derived from a Nash-bargaining, it should be bounded between 0 and 1. In this respect, the negative estimate of $\alpha_1$ is puzzling. However, it can be explained if the establishments' market share also affects wages in the absence of collective bargaining. This happens for example if high market share establishments act as monopsonies and hire a specialized workforce that cannot easily find a job in another firm. In this case, the market share should also enter in the set of covariates $Z_j$ that affect the market wage and $\alpha_1$ is no longer identifiable[25]. Finally, the difference in hourly wage between high market share and high bargaining power establishments in those with only a high market share is 1.75€, which is about 13% of the mean wage in the sample (12.69€). Columns (3) and (4) present the same estimations but using non-aggregated market share and bargaining power variables (the non-aggregated bargaining power variable is equal to 0 for non-unionized establishments and increases from 1 to 5 according to the percentage of unionized workers in establishments in which unions are recognized). The estimated cross-effect of market share and bargaining power is still positive and statistically significant. Finally models (1) to (4) have been reproduced using the log of the hourly wage as a dependent variable instead of the hourly wage. In the sixth column for example, the cross-effect on wages of a high bargaining power and a high market share is estimated to be around 8%.

## Selection issues and robustness of the results

Since predictions 1 and 2 are directly derived from a standard bargaining model, their empirical validation reinforces the rent-sharing interpretation of the union wage premium obtained by unions. But a selection process can still be at play, for example if the best workers

[25] In the same way, $\alpha_4$ is not identifiable since it is confounded with the constant term.

select themselves in large unionized firms which have the highest market share. Despite all my efforts to control for observable individual characteristics, some unobservable component of the individual productivity correlated with the establishments' union status and market share could lead to higher wages. This is unlikely for two reasons. First, the presence of unions in French firms relies on the individual willingness of some particular workers to accept to become union representatives. The union-status of firms depends more on a few individualities than on a global organization process of workers in a common group of interests. As a consequence, the union-status of a firm can be expected to be more exogenous with respect to the individual characteristics of the non-unionized workers who probably did not participate in the firm's unionization process. The surveyed employees in the REPONSE04 dataset are asked if they are unionized. Models (5) and (6) of table 5 reproduce the test of predictions 1 and 2 on the subsample of non-unionized workers only. The estimated effects on wages of the interaction between union recognition and market share and of the interaction between union recognition and unionization rate are only slightly lower than those found on the entire sample and significant at conventional statistical levels. Since the proxy I use for union bargaining power is the percentage of unionized workers and that unionized workers are likely to differ from non-unionized ones, the replication on non-unionized workers carried out in model (6) is even more crucial: it shows that the wage premium due to a higher bargaining power is similar for the subgroup of workers that induce this higher bargaining power than for the workers who legally benefit from the bargaining without being actively taking part in it. Overall, models (5) and (6) show that the empirical validation of predictions 1 and 2 is not driven by unionized workers that are far more likely to be selected. But an unexplainable selection process of the best non-unionized workers in establishments with both a high market share and a union recognized can still be at play and explain the results in model (5). Similarly, a selection process of the best non-unionized workers in establishments with both a high unionization rate and a union recognized can still be at play and explain the results in model (6). Since the point estimates for union recognition and market share (resp. unionization rate) in model (5) (resp. model 6) are all negative, such a selection process needs to take place simultaneously on both variables and cannot be a simple combination of two selection processes in establishments where unions are recognized and in establishments with a high market share (resp. high unionization rate). Also, if a selection process of the best workers is at play, this should show up in the labor productivity of the selected establishments. I have thus reproduced the empirical analysis of this paper including a linear control for labor productivity (measured as the value added per employee) in all the regression models in order to better capture worker's unobserved ability. The labor productivity variable I use comes from the DIANE dataset which contains publicly-available company accounts and that I have matched with the REPONSE04 dataset and is only defined at the firm level. Its inclusion in regression models tends in fact to increase both the point estimates and the significance of the estimates for the variable of interest of this study, suggesting that there is no selection of the best workers into organized establishments with a high market share or a high unionization rate[26].

Another potential source of biases is the selection by unions of the firms with the highest potential of rent extraction. In this case, the selection does not operate on the individual characteristics of the workers but on the characteristics of their working establishment. If unions target and organize the establishments with the highest potential rents, we should observe a correlation between union recognition and market share. The descriptive statistics in table 4 show that the distribution of unionized and non-unionized establishments across market share groups are very close. The link between market share and union recognition is tested more properly in table 7 which displays the results of 2 establishment-level logit regressions of union recognition on the market share indicator and a range of establishments' characteristics. In both models, a higher market share is not associated with a higher probability of union recognition. The marginal effect for the average establishment in the sample of switching from a market share category to the next one reported in the third row is close to 1%. This means for example that the differential in the probability of union recognition between the establishments with a market share lower than 5% (first group) and those with a market share larger than 50% (fourth group) is around 3%.

The market share variable used in this paper is a simple indicator of the presence of quasi-rents. If unions select firms, they should do so on the ground of their potential or actual rents. The fact that union presence and market share are not correlated does not necessarily mean that the unobserved component of quasi-rents $\mu_j$ is not correlated with union recognition. If such a selection is at play, one could expect that the unionized establishments declaring a low market share have in fact a higher $\mu_j$ than those declaring a high market share. This in turn would imply that the estimated relationship between market share and the

[26] Results not reported but available on demand. The DIANE dataset is provided by Bureau van Dijk, a private consulting company, and it is the French source file for the more famous Amadeus database. The match with REPONSE04 leads to a loss of about 500 REPONSE04 establishments, which is a quarter of the REPONSE sample. This is the reason why I have not de facto controlled for labor productivity in the regression models.

union wage gap plotted in figure 2 is flatter than it should be. In other words, the strategic selection of firms by unions according to their quasi-rents can only bias downward our estimates.

As discussed above, the union status of a firm is more driven by a few individualities than by a global organization process of workers. It is thus not surprising to find that the union status is not highly correlated with market share. However, the unionization rate variable is an aggregate measure of the workers' willingness to organize and should be more largely driven by the amount of rents available. Columns 3 and 4 of table 7 present establishment-level ordered-logit regressions of the unionization rate on the market share indicator and a range of establishments' characteristics. In the specification which includes detailed covariates and industry dummies (column 4), the effect of market share on the unionization rate is statistically significant at the 10% level. The fact that workers bargaining power and quasi-rents are positively correlated leads to a positive bias in the previous estimations of the link between union wage premiums and market share or unionization rates that would tend to make steeper the estimated relationships plotted in table 2. Since such a bias derives from a rent-extraction view of unions (there is no sense for a selection process of firms regarding their rents if workers cannot expect to bargain these rents), it actually cannot be advocated to prove that union wage premiums are not due to rent-extraction.

From a theoretical point of view, the parameter that is endogenous to the amount of quasi-rents available is the workers' bargaining power. Consider indeed that workers taken collectively have to pay a cost $c(\psi)$ to get a bargaining power equal to $\psi$, with $c(\psi)$ a convex increasing function. $c(\psi)$ is a simple measure of the aggregate cost paid at the establishment-level by workers to organize. It can for example comprise union dues for unionized workers, union work for union representatives and sunk costs invested in the organization process[27]. In a two-step game, the workers should choose in the first step the bargaining power that will maximize their surplus in the subsequent bargaining with the employer:

$$\psi_j = Arg\max(w_j^m + \psi_j QR_j - c(\psi_j)$$

The chosen bargaining power simply satisfies $c'(\psi_j) = QR_j$ and is an increasing function of the expected quasi-rents. I test this link between workers' bargaining power and quasi-rents by estimating a logit regression of the high bargaining power indicator (that combines together union recognition and unionization rate) on the establishments' market share and

[27] The standard question of which workers in particular bear this cost and which are free-riders is beyond the scope of this paper and is not modeled in the aggregate establishment-level cost $c(\psi)$.

other covariates. The results are displayed in columns 5 and 6 of table 7: jumping from a market share group to the next has a positive and statistically very significant impact on the probability to have high bargaining power. Going from a market share lower than 3% to a market share larger than 50% increases the likelihood of the average establishment in the sample to have a high bargaining power by about 10%. More generally, the endogeneity of the workers' bargaining power induces an upward bias in studies that attempt to capture the average degree of rent-sharing with regressions of individual wages on measures of quasi-rents (this is potentially the case for the studies by Abowd and Lemieux 1993, Abowd and Allain 1995, Blanchflower et al 1996 and Kramarz 2008).

A last alternative to the rent-sharing explanation of the union wage premium is the theory of compensating wage differentials. If workers can move with no cost from the non-unionized to the unionized sector, market forces should make them indifferent in equilibrium between working in one or the other of these two sectors. In this case, the wage premium obtained by unions should be compensated by losses on non-wage aspects of workers' compensation such as working effort, working conditions or job protection. If unions' objective function favors wages at the expense of non-wage aspects of compensation, unions could in fact extract monetary rents in firms where they are present and lose on other aspects in the same time. The simple bargaining model I use in this paper does not model the non-wage aspects of workers' compensation and makes it impossible to sort out between this weaker explanation and the pure rent-extraction story that would predict that unions manage to make workers better off on all the aspects of their compensation. A key prediction of the compensating wage differential theory is that workers should remain at the equilibrium indifferent between working in the unionized and the non-unionized sectors. I test this prediction in appendix C by looking at voluntary quits in unionized and non-unionized firms. I show that the annual rate of voluntary quits is lower by one third in unionized establishments, even when I control for workers' productivity. This is consistent with the idea that workers might be better off in these establishments. Furthermore, the quarterly dismissal rate is also a bit lower in these establishments, suggesting that job protection is not worse when unions are present. These two results suggest that the wage premium obtained by unions is not compensated by losses on other aspects, contrary to what is predicted by the theory of compensating wage differentials[28] .

[28] Looking at workers' moves and job protection can be computationally intensive. Details on the dataset I use of and on my empirical specifications are provided in appendix C.

## Conclusion

This paper studies the wage premium associated with establishment-level union recognition in France. A premium of 2 to 3% is precisely estimated in individual hourly wage equations using a large dataset that enables to control for standard observable workers and firms characteristics. Despite the reputation of French unions as being strong and the fact that the system of French industrial relations is to a large extent decentralized, this premium is far lower that what is found in the literature with similar techniques for other countries (Card and De La Rica, 2006). I suggest that this is due to the weak legal barriers to firm-level union recognition in France, which explains both that unions are recognized in a large number of firms but with a low bargaining power on average in these firms.

This paper then pushes the analysis further and investigates whether the union wage premium in France is due to rent-extraction. If this is the case, the premium should be increasing with the amount of rents available in firms and with the bargaining power of unions. I derive these two predictions from a simple bargaining model and test it using two proxy variables for the existence of potential rents and the unions' bargaining power. Empirical results show that the union wage premium increases from virtually 0 to 8% to 12% in firms with high potential rents and when unions have a strong bargaining power. This confirms the theoretical predictions of the bargaining framework.

The paper also contributes to the rent-sharing literature along two lines. First, it shows that the workers' bargaining power is likely to increase with the amount of rents they can get. This implies that usual estimates of this rent-sharing parameter based on regressions of the individual wage on measures of quasi-rents might be partially biased. Second, it is the first in the literature to estimate a bargaining model with proxy variables for both quasi-rents and for the workers' bargaining power.

Finally, the union wage premium I find in firms with high potential rents is in line with what is usually found in the international literature for all firms. This suggests that in countries where the cost of organizing is higher than in France, unions organize a smaller number of firms and target only those in which the amount of rents they can extract will lead to a wage premium that is sufficient to compensate this higher shadow cost to organize the firm. Further research would enable to link more precisely the legal cost paid by unions to organize to the wage premium they obtain in each country.

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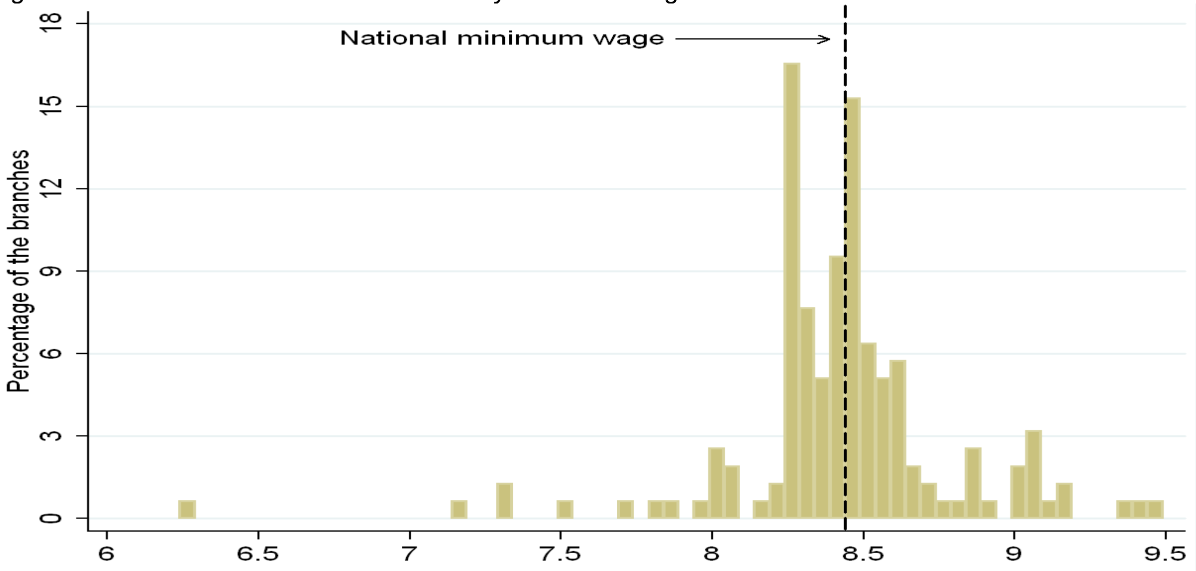
## Tables and Figures

*Figure 1.* Distribution of the branches hourly minimum wages in 2007.




*Notes:* The hourly national minimum wage in the 1st of July 2007 was equal to 8.44€. *Source:* Author's computations from the Collective Bargaining annual book (2006). Only the 160 branches covering more than 5,000 workers are represented. They cover 10 million workers, which is more than one half of the private sector total employment.

*Table 1.* Description of French main unions:

| **French union** | **Historical/ Ideological roots** | **Representation (in 2004)** |
|---|---|---|
| Confédération Générale du Travail (CGT) | Marxism | 27.6% |
| Confédération Française Démocratique du Travail (CFDT) | Socialism | 27.3% |
| Force Ouvrière (FO) | Trotskyism | 19.7% |
| Confédération Générale des Cadres (CGC) | white collars | 11.3% |
| Confédération Française des Travailleurs Chrétiens (CFTC) | Christians | 10.5% |
| Others (these are generally local or sector specific unions) | | 3.5% |

*Notes:* The last column gives the distribution of all the union delegates among establishments with more than 20 employees in 2004 (obtained from the REPONSE dataset using a weighted average of the number of union delegates in each workplace).

*Table 2.* Distribution of the workplaces with more than 20 employees in terms of the total number of unions present

| Number of unions present in a workplace | Proportion of workplaces in the REPONSE data | Proportion of French workplaces concerned | Proportion of French workers concerned |
|---|---|---|---|
| 0 | 33.90% | 64.34% | 36.03% |
| 1 | 18.09% | 19.32% | 19.04% |
| 2 | 13.29% | 7.60% | 13.07% |
| 3 | 12.16% | 3.73% | 10.44% |
| 4 | 9.98% | 2.60% | 8.37% |
| 5 | 9.38% | 1.98% | 8.78% |
| 6 | 2.36% | 0.27% | 3.34% |
| more than 6 | 0.83% | 0.17% | 0.94% |

*Note:.* From Author's computation using the REPONSE dataset and the set of weights provided by the ministry of Labor to make the data sample representative of the French private sector workplaces with more than 20 employees or of the workers in those workplaces.

*Table 3.* Log Hourly Wage Regressions (ESS02)

| | *Dependent variable: log of gross hourly wage (from ESS02)* | | | | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Workplace Union Recognition | 0.201 | 0.0272 | 0.0173 | | | |
| | (0.007) | (0.005) | (0.005) | | | |
| Firm Wage Contract | | | | 0.073 | -0.006 | -0.007 |
| | | | | (0.015) | (0.008) | (0.007) |
| *Worker's characteristics* | | | | | | |
| Women | | -0.132 | -0.133 | | -0.133 | -0.136 |
| | | (0.003) | (0.003) | | (0.005) | (0.005) |
| High School | | 0.097 | 0.099 | | 0.099 | 0.101 |
| | | (0.004) | (0.004) | | (0.005) | (0.005) |
| Some College | | 0.136 | 0.139 | | 0.141 | 0.143 |
| | | (0.004) | (0.004) | | (0.006) | (0.006) |
| College or University Degree | | 0.268 | 0.277 | | 0.274 | 0.282 |
| | | (0.005) | (0.005) | | (0.008) | (0.008) |
| Age (in years) | | 0.011 | detailed | | 0.012 | detailed |
| | | (0.0001) | | | (0.0002) | |
| Fixed Term Contract | | -0.055 | -0.021 | | -0.052 | -0.017 |
| | | (0.008) | (0.008) | | (0.011) | (0.012) |
| *Firm's characteristics (reference:workplaces with 10 to 20 workers)* | | | | | | |
| 20-50 Workers | | 0.030 | 0.024 | | 0.018 | 0.013 |
| | | (0.006) | (0.026) | | (0.009) | (0.009) |
| 51-100 Workers | | 0.036 | 0.028 | | 0.032 | 0.024 |
| | | (0.007) | (0.026) | | (0.010) | (0.009) |
| 101-200 Workers | | 0.050 | 0.043 | | 0.056 | 0.047 |
| | | (0.007) | (0.026) | | (0.009) | (0.009) |
| Over 200 Workers | | 0.075 | 0.067 | | 0.083 | 0.075 |
| | | (0.007) | (0.026) | | (0.009) | (0.009) |
| Intercept | 2.63 | 2.71 | 2.87 | 2.73 | 2.69 | 2.86 |
| | (0.005) | (0.023) | (0.027) | (0.012) | (0.037) | (0.020) |
| Industries | No | 1 digit | 2 digits | No | 1 digit | 2 digits |
| Observations | 106,734 | 97,751 | 97,751 | 70,987 | 64,987 | 64,987 |
| R-squared | 0.034 | 0.625 | 0.636 | 0.000 | 0.626 | 0.637 |

*Notes:* All models except (1) and (4) also include 10 indicators for region and 4 indicators for occupation. Standard errors are calculated with clustering by establishments in models (1) to (3) and by firms in models (4) to (6). Models (2) and (5) include 9 indicators for industry. Models (3) and (6) include 47 indicators for industry as well as 10 indicators for worker's age and 4 indicators for worker tenure.

*Table 4.* Distribution of establishments in terms of their declared market share, targeted market and percentage of unionized workers (REPONSE04)

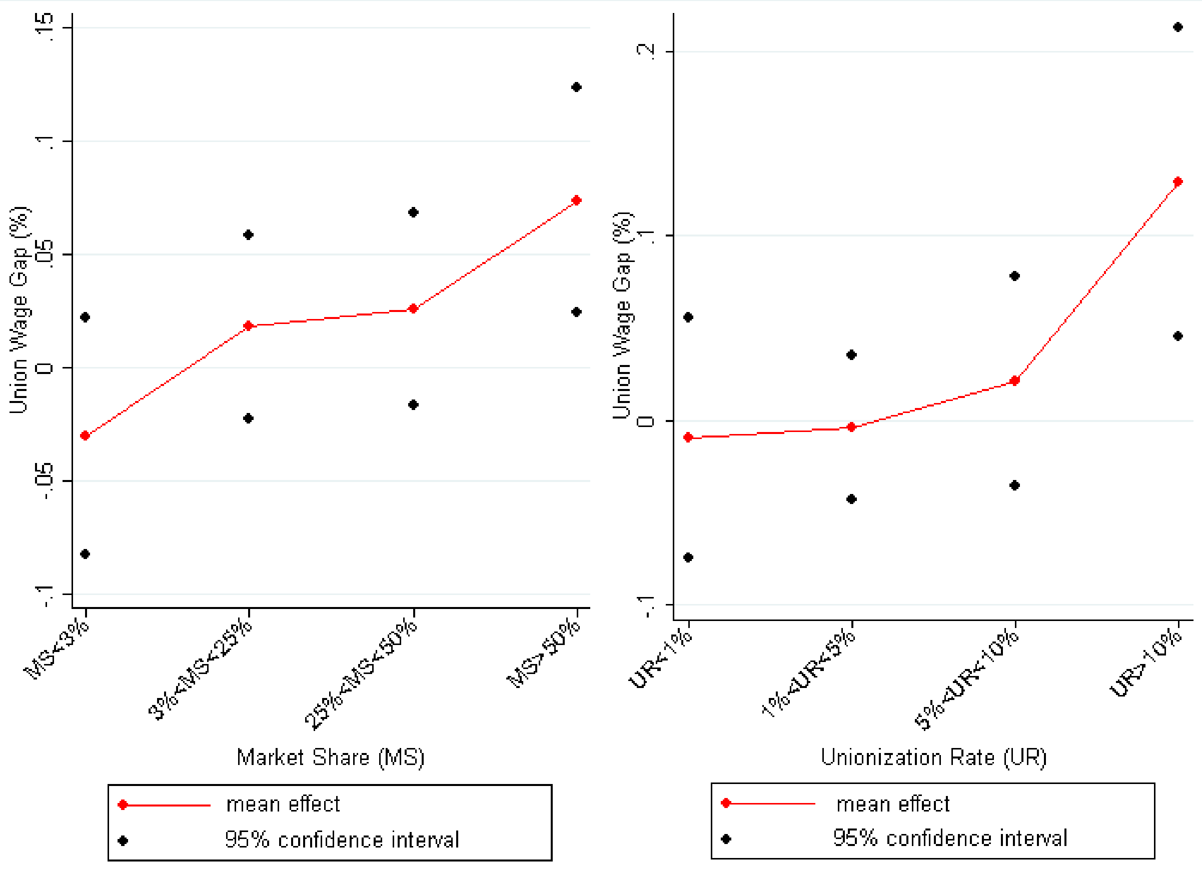
| **Market Share (MS)** | **MS<3%** | **3%<MS<25%** | **25%<MS<50%** | **MS>50%** | | **Total** |
|---|---|---|---|---|---|---|
| Number of establishments | 318 | 787 | 451 | 305 | | 1861 |
| (percentage) | (17%) | (42%) | (24%) | (18%) | | (100%) |
| Nb. of non-unionized estab. | 146 | 257 | 137 | 109 | | 649 |
| (percentage) | (22%) | (40%) | (21%) | (17%) | | (100%) |
| Nb. of unionized estab. | 172 | 530 | 314 | 196 | | 1212 |
| (percentage) | (14%) | (44%) | (26%) | (16%) | | (100%) |
| **Targeted market** | **Local** | **Regional** | **National** | **European** | **International** | **Total** |
| Number of establishments | 477 | 416 | 576 | 305 | 666 | 2440 |
| (percentage) | (20%) | (17%) | (24%) | (12%) | (27%) | (100%) |
| **Unionization Rate (UR)** | **UR<1%** | **1%<UR<5%** | **5%<UR<10%** | **UR>10%** | | **Total** |
| Number of establishments | 493 | 560 | 595 | 481 | | 2129 |
| (percentage) | (23%) | (26%) | (28%) | (23%) | | (100%) |
| Nb. of non-unionized estab. | 473 | 190 | 72 | 18 | | 753 |
| (percentage) | (63%) | (25%) | (9.6%) | (2.4%) | | (100%) |
| Nb. of unionized estab. | 20 | 370 | 523 | 463 | | 1376 |
| (percentage) | (1.45%) | (27%) | (38%) | (34%) | | (100%) |

*Table 5*. Log Hourly Wage Regressions: Union recognition, bargaining power and rents (REPONSE04)

| | All surveyed workers | | | | Non-unionized only | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Union Recognition (establishment level) | -0.028<br>(0.030) | -0.046<br>(0.030) | -0.025<br>(0.020) | -0.018<br>(0.019) | -0.039<br>(0.032) | -0.014<br>(0.020) |
| Market share (4 groups) | -0.018<br>(0.008) | -0.022<br>(0.008) | | | -0.024<br>(0.008) | |
| Union Recognition*Market share | 0.027<br>(0.011) | 0.028<br>(0.011) | | | 0.025<br>(0.011) | |
| Unionization Rate (5 groups) | | | -0.010<br>(0.010) | -0.012<br>(0.009) | | -0.009<br>(0.010) |
| Union Recognition*Unionization rate | | | 0.029<br>(0.012) | 0.024<br>(0.011) | | 0.020<br>(0.012) |
| *Workers controls:* | | | | | | |
| Gender, educ., age, occup., full time | Yes | Yes | Yes | Yes | Yes | Yes |
| Detailed age and tenure | No | Yes | No | Yes | No | Yes |
| *Establishment controls:* Size, Region | Yes | Yes | Yes | Yes | Yes | Yes |
| *Establishment controls:* Age | No | Yes | No | Yes | No | Yes |
| *Establishment controls:* Industries | 1 digit | 3 digits | 1 digit | 3 digits | 3 digits | 3 digits |
| Observations | 4990 | 4934 | 5612 | 5547 | 4310 | 4841 |
| R-squared | 0.641 | 0.711 | 0.640 | 0.714 | 0.695 | 0.725 |

*Notes:* In all models, union recognition, market share and unionization rate are defined at the establishment level and standard errors are calculated with clustering by establishments. The control variables used are those in columns (2) and (3) of table A3. Models (5) and (6) are estimated on the subsample of non-unionized workers.

*Figure 2.* The union wage gap in each market share and unionization rate group (controlling for other observable characteristics)




*Notes:* The union wage gaps in both graphs are obtained by running 2 regressions of the individual hourly wage from DADS03 on detailed observable individual characteristics (gender, age tenure, education, occupation, full-time job) and establishment characteristics (size, region, firm age, 3-digit industries) and a set of 4 indicators for market share groups (left graph) or 5 indicators for unionization rate group (right graph) as well as the interaction of these indicators with a union recognition dummy. The plotted point estimates and 95% confidence intervals correspond to the estimated effect of these interactions on hourly wages. The point estimates should be interpreted as the union wage premium within each market share or unionization rate group, conditional on other observable workers and establishment characteristics.

*Table 6*. Estimation of structural wage and log-wage equations with proxy variables for quasi-rents and bargaining power (REPONSE04)

| *Dependent variable:* | *Net hourly wage* | | | | *Log of Net hourly wage* | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (1) | (2) | (3) | (4) |
| High Market Share (HMS) | -0.57 | -0.73 | | | -0.031 | -0.039 | | |
| | (0.24) | (0.26) | | | (0.015) | (0.015) | | |
| High Bargaining Power (HBP) | -0.02 | -0.03 | | | 0.023 | 0.013 | | |
| | (0.25) | (0.22) | | | (0.014) | (0.013) | | |
| HMS*HBP | 1.42 | 1.78 | | | 0.067 | 0.083 | | |
| | (0.85) | (0.87) | | | (0.039) | (0.037) | | |
| Market Share (MS, 4 groups) | | | -0.36 | -0.48 | | | -0.019 | -0.022 |
| | | | (0.17) | (0.18) | | | (0.008) | (0.008) |
| Bargaining Power (BP, 5 groups) | | | -0.30 | -0.43 | | | -0.004 | -0.009 |
| | | | (0.24) | (0.25) | | | (0.009) | (0.009) |
| MS*BP | | | 0.13 | 0.18 | | | 0.007 | 0.007 |
| | | | (0.07) | (0.08) | | | (0.003) | (0.003) |
| *Workers controls:* | | | | | | | | |
| Gender educ. age occup. full time | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Detailed age and tenure | No | Yes | No | Yes | No | Yes | No | Yes |
| *Establishment controls:* Size Region | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| *Establishment controls:* Age | No | Yes | No | Yes | No | Yes | No | Yes |
| Industries | 1 digit | 3 digits | 1 digit | 3 digits | 1 digit | 3 digits | 1 digit | 3 digits |
| Observations | 4430 | 4386 | 4430 | 4386 | 4430 | 4386 | 4430 | 4386 |
| R-squared | 0.393 | 0.425 | 0.393 | 0.425 | 0.633 | 0.684 | 0.633 | 0.684 |

*Notes:* In all models, standard errors are calculated with clustering by establishments. The high market share variable is a dummy equal to 1 for establishments declaring a market share larger than 50%. The high bargaining power variable is a dummy equal to 1 for establishments where unions are recognized and with more than 10% unionized workers. The bargaining power variables is equal to 0 for establishments in which unions are not recognized and varies from 1 to 5 according to the proportion of unionized workers for establishments in which they are recognized.

*Table 7*. Establishment-level regressions: Are Union recognition and bargaining power explained by rents? (REPONSE04)

| *Dependent variable:* | *Union Recognition* | | *Unionization Rate* | | *High Bargaining Power* | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Market share (4 groups) | 0.051 | 0.059 | 0.104 | 0.095 | 0.183 | 0.195 |
| (robust standard error) | (0.065) | (0.072) | (0.049) | (0.054) | (0.066) | (0.076) |
| *Marginal effect at the mean* | *0.010* | *0.013* | - | - | *0.028* | *0.030* |
| *Establishment controls:* | | | | | | |
| Size and region | Yes | Yes | Yes | Yes | Yes | Yes |
| Age | No | Yes | No | Yes | No | Yes |
| Industries | 1 digit | 3 digits | 1 digit | 3 digits | 3 digits | 3 digits |
| Regression Model* | Logit | Logit | Ologit | Ologit | Logit | Logit |
| Observations | 1860 | 1591 | 1647 | 1644 | 1639 | 1466 |
| Pseudo R-squared | 0.365 | 0.367 | 0.113 | 0.160 | 0.126 | 0.182 |

*Notes:* The market share variable takes values 1 to 4. The high bargaining power variable used in the last 2 columns is a dummy equal to 1 for establishments where unions are recognized and with more than 10% unionized workers. The control variables used are those in columns (2) and (3) of table A3.
* "Ologit" means "ordered logit" model.

## Appendix A: Description of ESS02 and REPONSE04 variables

*Table A1*. Means of individual and establishment-level variables in ESS02

| **Individual variables:** | Whole Sample (97,751 individuals) | | Unionized Establishments (57,435 individuals) | | Non-Unionized Establishments (43,316 individuals) | |
|---|---|---|---|---|---|---|
| | Mean | Std. Dev. | Mean | Std. Dev. | Mean | Std. Dev. |
| Gross hourly wage (€) | 15.09 | 9.26 | 16.14 | 9.56 | 13.75 | 8.67 |
| Women | 0.35 | 0.48 | 0.33 | 0.47 | 0.37 | 0.48 |
| Age | 39.15 | 10.51 | 40.21 | 10.32 | 37.79 | 10.60 |
| Tenure | 11.32 | 10.25 | 13.48 | 10.89 | 8.53 | 8.59 |
| Long-term contract | 0.93 | 0.25 | 0.94 | 0.24 | 0.93 | 0.26 |
| *Education:* | | | | | | |
| Less than high school | 0.59 | 0.49 | 0.58 | 0.49 | 0.61 | 0.49 |
| High school degree | 0.17 | 0.37 | 0.16 | 0.37 | 0.17 | 0.38 |
| Some College | 0.13 | 0.34 | 0.13 | 0.34 | 0.13 | 0.34 |
| College or University Degree | 0.11 | 0.31 | 0.12 | 0.33 | 0.09 | 0.28 |
| *Occupation:* | | | | | | |
| Blue Collar | 0.40 | 0.49 | 0.39 | 0.49 | 0.40 | 0.49 |
| Clerk | 0.20 | 0.40 | 0.17 | 0.38 | 0.24 | 0.43 |
| Supervisor or Technician | 0.24 | 0.43 | 0.25 | 0.43 | 0.22 | 0.41 |
| Manager | 0.17 | 0.37 | 0.18 | 0.39 | 0.14 | 0.35 |
| **Establishment variables:** | Whole Sample (10,741 estab.) | | Unionized Establishments (5,659 estab.) | | Non-Unionized Establishments (5,082 estab.) | |
| | Mean | Std. Dev. | Mean | Std. Dev. | Mean | Std. Dev. |
| Union recognition | 0.22 | 0.41 | 1.00 | 0.00 | 0.00 | 0.00 |
| Size | 59.07 | 1496.43 | 179.57 | 3047.14 | 25.84 | 540.11 |

*Notes:* Individual (resp. establishment) variables are weighted by ESS02 workers (resp. establishments) sampling weights.

*Table A2*. Means of individual and establishment-level variables in REPONSE04

| **Individual variables:** | Whole Sample (6,629 individuals) | | Unionized Establishments (4,459 individuals) | | Non-Unionized Establishments (2,152 individuals) | |
|---|---|---|---|---|---|---|
| | Mean | Std. Dev. | Mean | Std. Dev. | Mean | Std. Dev. |
| Net hourly wage (€) | 12.60 | 8.23 | 13.36 | 8.01 | 11.18 | 8.47 |
| Women | 0.38 | 0.49 | 0.37 | 0.48 | 0.41 | 0.49 |
| Age | 39.95 | 9.80 | 40.65 | 9.73 | 38.65 | 9.83 |
| Tenure | 11.77 | 10.17 | 13.25 | 10.58 | 9.01 | 8.75 |
| Full-time worker | 0.90 | 0.29 | 0.91 | 0.29 | 0.89 | 0.31 |
| *Education:* | | | | | | |
| Less than high school | 0.57 | 0.50 | 0.55 | 0.50 | 0.60 | 0.49 |
| High school degree | 0.13 | 0.34 | 0.13 | 0.34 | 0.14 | 0.35 |
| Some College | 0.15 | 0.35 | 0.15 | 0.36 | 0.14 | 0.35 |
| College or University Degree | 0.15 | 0.36 | 0.17 | 0.37 | 0.12 | 0.32 |
| *Occupation:* | | | | | | |
| Blue Collar | 0.36 | 0.48 | 0.36 | 0.48 | 0.36 | 0.48 |
| Clerk | 0.17 | 0.37 | 0.14 | 0.35 | 0.22 | 0.41 |
| Supervisor or Technician | 0.27 | 0.45 | 0.28 | 0.45 | 0.26 | 0.44 |
| Manager | 0.20 | 0.40 | 0.22 | 0.42 | 0.16 | 0.37 |

| **Establishment variables:** | Whole Sample (2,451 estab.) | | Unionized Establishments (1,612 estab.) | | Non-Unionized Establishments (839 estab.) | |
|---|---|---|---|---|---|---|
| | Mean | Std. Dev. | Mean | Std. Dev. | Mean | Std. Dev. |
| Union recognition | 0.36 | 0.48 | 1.00 | 0.00 | 0.00 | 0.00 |
| Size | 77.24 | 172.20 | 134.16 | 270.37 | 45.29 | 39.37 |
| Establishment Age: | | | | | | |
| less than 5 years | 0.04 | 0.20 | 0.04 | 0.21 | 0.04 | 0.20 |
| 5 to 9 years | 0.10 | 0.30 | 0.08 | 0.27 | 0.11 | 0.31 |
| 10 to 19 years | 0.28 | 0.45 | 0.26 | 0.44 | 0.29 | 0.45 |
| 20 to 49 years | 0.40 | 0.49 | 0.39 | 0.49 | 0.41 | 0.49 |
| 50 years or more | 0.18 | 0.38 | 0.23 | 0.42 | 0.15 | 0.36 |
| ICT use | 7.19 | 3.44 | 7.65 | 3.20 | 6.95 | 3.55 |
| Managerial practices | 4.65 | 1.60 | 5.03 | 1.47 | 4.43 | 1.63 |
| Belong to a listed firm | 0.02 | 0.12 | 0.02 | 0.15 | 0.01 | 0.11 |
| Belong to a family Firm | 0.59 | 0.49 | 0.42 | 0.49 | 0.69 | 0.46 |

*Notes:* Individual (resp. establishment) variables are weighted by ESS02 workers (resp. establishments) sampling weights. ICT use and managerial practices are aggregated indexes.

*Table A3*. Log Hourly Wage Regressions (REPONSE04)

| | *Dependent variable: log of net hourly wage* | | | | |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| Workplace Union Recognition | 0.191 | 0.031 | 0.019 | 0.018 | 0.021 |
| | (0.016) | (0.013) | (0.013) | (0.013) | (0.014) |
| *Worker's characteristics* | | | | | |
| Women | | -0.139 | -0.134 | -0.135 | -0.145 |
| | | (0.008) | (0.008) | (0.008) | (0.009) |
| High School | | 0.056 | 0.056 | 0.048 | 0.054 |
| | | (0.010) | (0.009) | (0.009) | (0.011) |
| Some College | | 0.111 | 0.111 | 0.108 | 0.115 |
| | | (0.012) | (0.012) | (0.011) | (0.013) |
| College or University Degree | | 0.222 | 0.235 | 0.224 | 0.219 |
| | | (0.017) | (0.017) | (0.016) | (0.018) |
| Age | | 0.011 | detailed | detailed | detailed |
| | | (0.000) | | | |
| Full time | | 0.068 | 0.051 | 0.038 | 0.027 |
| | | (0.012) | (0.012) | (0.012) | (0.014) |
| *Firm's characteristics* *(reference in (5): workplaces with 20 to 50 workers, not intensive in ICT, not family and not listed)* | | | | | |
| 51-100 Workers | | 0.000 | 0.000 | -0.003 | -0.004 |
| | | (0.015) | (0.014) | (0.015) | (0.017) |
| 101-200 Workers | | 0.028 | 0.030 | 0.028 | 0.006 |
| | | (0.016) | (0.015) | (0.015) | (0.018) |
| Over 200 Workers | | 0.074 | 0.060 | 0.064 | 0.049 |
| | | (0.015) | (0.015) | (0.015) | (0.019) |
| ICT use | | | | | 0.009 |
| | | | | | (0.002) |
| Managerial practices | | | | | 0.007 |
| | | | | | (0.004) |
| Listed | | | | | 0.081 |
| | | | | | (0.035) |
| Family firm | | | | | -0.023 |
| | | | | | (0.011) |
| Intercept | 2.300 | 2.365 | 2.495 | 2.475 | 2.385 |
| | (0.013) | (0.041) | (0.038) | (0.041) | (0.054) |
| Industries | No | 1 digit | 2 digits | 3 digits | 4 digits |
| Observations | 6610 | 6449 | 6357 | 6357 | 4990 |
| R-squared | 0.042 | 0.628 | 0.656 | 0.678 | 0.709 |

Notes: All models except (1) also include 10 indicators for region and 4 indicators for occupation. Standard errors are calculated with clustering by establishments in all models. Model (2) includes 14 indicators for industry. Model (3), (4) and (5) include 10 indicators for worker's age, 4 indicators for worker's tenure and 5 indicators for firm's age. Models (3), (4) and (5) include respectively 51, 168 and 328 indicators for industry.

Tables A1 and A2 show that workers in unionized establishments are less often women and are older with a longer tenure. The education and occupation profiles of workers in unionized and non-unionized establishments are close with only slightly less clerks in unionized establishments. Unionized workplaces are a lot larger. Table A2 shows that they are also older and more intensive in ICT. Finally, they have more innovative managerial practices and belong to a family less often.

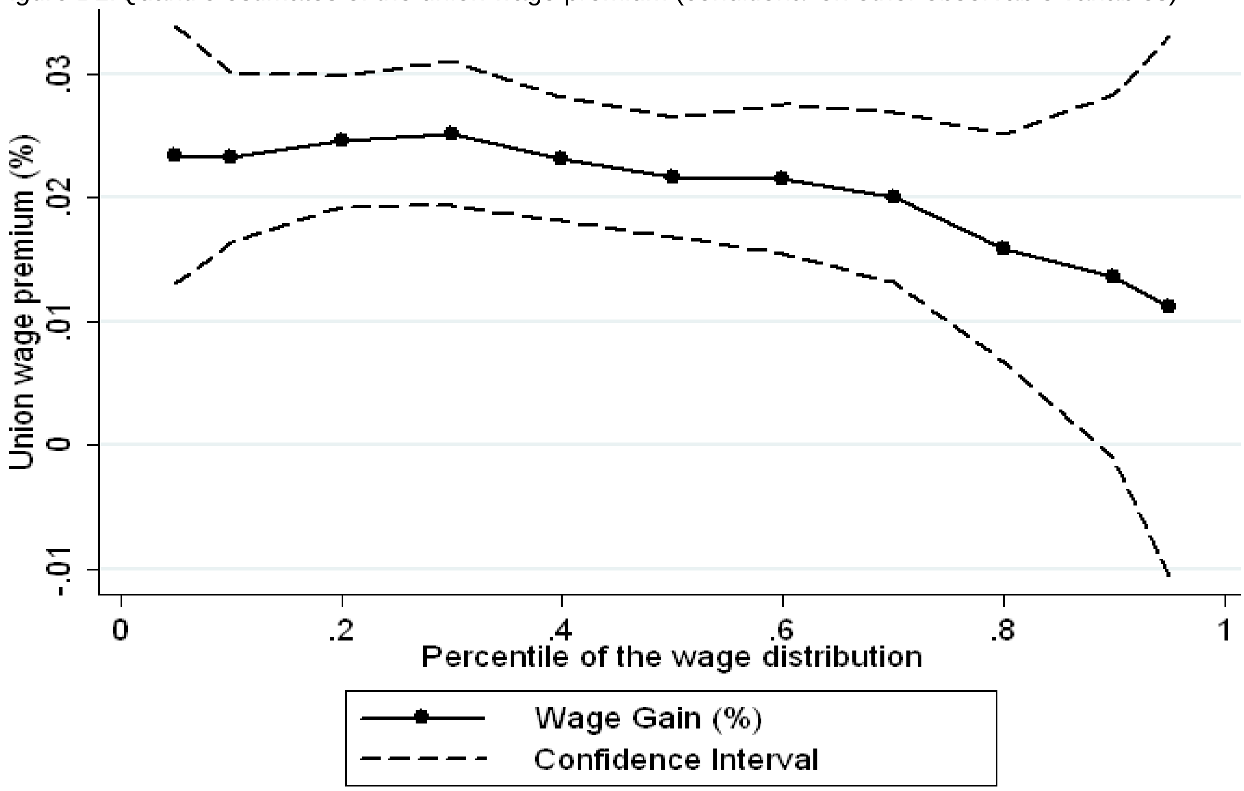
Models (1), (2) and (3) of table A3 obtained with REPONSE04 are equivalent to models (1), (2) and (3) of table 3 obtained with ESS02 with the exception that the type of working contract is not observable in REPONSE04 and has been replace by a dummy variable for full time workers (which is observable only in REPONSE04). The estimated wage premium associated with union recognition is very close in both tables. Standard errors for all coefficients in table A3 are about twice larger than those in table 3. This is coherent with the fact that there are around 3000 establishments of firms in REPONSE and around 13000 in ESS02 (since standard errors decrease with the square root of the number of observations, a multiplication by 4 of the sample size indeed corresponds to a division by 2 of the estimated standard errors) but renders the estimated effect of the union wage premium insignificant at the 5% level in models (3), (4) and (5).

Model (4) of table A3 uses 4-digit instead of 3-digit industry controls. Model (5) uses the large amount of information available in REPONSE04 to control for other firm characteristics. First, unions play a role in their firm's decision to invest in Information and Communication Technology (ICT). An indicator of ICT intensity similar to the one used by Bassanini et al (2010) has thus been added in the regression model. Unions also influence management practices (being against performance pay for example). For this reason, an indicator of so-called modern management practices has been included in the model. Finally, I also include 2 control variables for listed and family firms since unions are known to be more present in large listed firms that offer larger wages and less present in family firms that offer lower wages (Muller Philipon 2006, Sraer Thesmar 2007).

## Appendix B: Unions and the structure of wages (test of prediction 3).

The more concise way to test prediction 3 (the larger a worker's wage, the smaller his log-wage premium) is probably to use quantile regression models. This is done in figure B1 which shows how union recognition shifts each quantile of the overall wage distribution, conditional on other observable variables. Coherent with prediction 3, the union wage premium decreases along the wage distribution. Union recognition is associated with an increase of the 3 first deciles of the wage distribution by roughly 2.5% whereas the shift for the 2 last deciles is only around 1.5%.

*Figure B1.*Quantile estimates of the union wage premium (conditional on other observable variables)




*Notes*. Obtained from a quantile regression that includes the detailed controls in model (3) of table 3. The individual controls are 10 workers' age groups, 4 workers' tenure groups as well as indicators for workers' gender, education, occupation and an indicator for full-time jobs. The establishment-level controls are 5 indicators for establishment size, 10 indicators for regions and 47 indicators for industries.

## Appendix C: Unions, voluntary quits and job protection (test of the absence of compensating wage differentials)

I investigate whether unionized firms offer a specific compensation package including more job security. In order to do so, I estimate the relationship between workplace-level union recognition and different types of separations rates. In my data, separation rates are available for each quarter over 2002-2006 whereas union recognition, establishment and firm-level controls are available only for year 2004. Some types of separations, including dismissals, fluctuate quite a lot over time. Therefore, I estimate them over a rather long time period (2002-2006). The model that I estimate is thus:

$$S^a_{jt} = \gamma U_j + X_j \beta + D_t + \varepsilon_{jt}$$

where $S^a_{jt}$ is the separation rate of type *a*, in establishment *j* at quarter *t*, $U_j$ is a dummy variable indicating the presence of one or more unions in establishment *j*, $X_j$ is a vector of establishment and firm-level controls and $D_t$ is a time dummy. The various types of separations include: dismissals, voluntary quits, retirement, end of trial period and end of fixed-term contract.

I use the DMMO/ EMMO database. The DMMO (Données sur les Mouvements de Main-d'Oeuvre) has exhaustive data on gross worker flows (hirings and separations, excluding workers provided by temporary help supply firms) for establishments with 50 employees or more for each quarter. The data is broken down by type of flow. The EMMO (Enquête sur les Mouvements de Main-d'Oeuvre) has identical information on a representative sample of establishments with less than 50 employees. I match these dataset with REPONSE 2004. The match of the two datasets is rather good: after doing it, I obtain 2024 matched establishments reporting information on unions. I use the DMMO-EMMO data to compute indicators of job security and, more specifically, of separation rates at the establishment level. In order to do that, I drop all movements corresponding to job spells shorter than one month. These indeed correspond to very short trial periods or temporary contracts which have little to do with job security for core workers[29]. I also exclude movements due to transfers between two establishments of the same firm. The data allows to build separation rates for each quarter over 2002-2006[30]. As standard in the gross worker flow literature (Davis et al, 2006), separation rates are defined as the sum of all types of exits[31] divided by average employment. In order to go deeper into the types of separations, I define dismissal rates, quit rates, retirement rates, rates of end of trial periods and rates of end of fixed-term contracts as the ratio of the corresponding type of exit during the quarter to the average employment of the quarter.

Each quarter, the establishments in my sample separate from about 4% of their workers. Organized establishments appear to have a lower separation rate than their non-organized counterparts (3.6% on average, against 5.3% in non-organized establishments). Each quarter, organized establishments dismiss fewer workers (0.47% on average, compared 0.60% in non-organized establishments) and have more separations due to retirement or end of trial period (0.18% and 0.13%, respectively, against 0.13% and 0.08% in non-organized establishments).

[29] The results are nonetheless robust to the inclusion of these very short job spells.

[30] Potentially we have information for 20 quarters for each establishment. However, there are several missing values and we have complete information for only 750 establishments. The average number of quarters with non-missing data per establishment is 16.3. We have non-missing data in at least half of the quarters in 1756 establishments, while we have less than one fourth of the quarters for 117 establishments.

[31] Exits can be due to dismissals, quits, retirement and early retirement, end of trial periods, end of fixed-term contracts or other temporary contracts, military service, injuries, death or unknown exits.

Less workers quit voluntarily (0.7% compared to 1.55% in non-organized establishments) or separate at the end of a fixed-term contract (1.9%, compared to 2.4% in non-organized establishments).
Table C1 shows that dismissal rates are a bit lower in establishments where unions are recognized even after controlling for a basic set of establishment and worker controls[32], including 3-digit industry dummies and a full set of time dummies. However, voluntary quits appear to be a lot lower (lower by one third) in organized workplaces, this difference being statistically very significant. This suggests that workers could be better off in organized workplaces and less willing to leave. It could also mean that they are less productive and have therefore fewer outside options. However, when I add workers productivity at the firm level as an additional control variable in the regression model, I still find that workers in organized workplaces leave their firm less often (the point estimate is -0.21, which is slightly slower than what is found in column 2 of table C1, and the standard error is 0.10). This additional result reinforces the first interpretation: workers in organized workplaces might leave less often because they are better off. This would in turn mean that the wage premium obtained by unions is not compensated by losses on other aspects.
Organized firms do not differ from their non-organized counterparts concerning retirements' rates. The retirements' rates are largely exogenous to the firms' managing choices and mostly driven by the age profiles of the workers. Since I control in the regression models for the workers characteristics, including their age, it is normal to find no difference in term of retirements' rates between organized and non-organized workplaces. The lower separation rate for core workers in organized establishments (e.g. those under open-ended contract) seems to be compensated by a larger separation rate for workers under fixed-term contract (even though this result is not statistically significant). This is consistent with the insider/outsider literature and the idea that unions defend at first core workers. Finally, there is no significant difference between organized and non-organized establishments when we consider all separations together (last column of table C1).

*Table C1*. Establishment-level regressions: Union recognition and separations (REPONSE04+DMMO)

| | *Dependent variable (quarterly rates):* | | | | | |
|---|---|---|---|---|---|---|
| | Dismissals | Quits | Retirem. | End-trial | End-temp. | All separat. |
| Union Recognition (establishment level) | -0.087* | -0.278*** | -0.007 | -0.135** | 0.254 | -0.239 |
| | (0.048) | (0.089) | (0.011) | (0.065) | (0.209) | (0.300) |
| time dummies | yes | yes | yes | yes | yes | yes |
| *Workers' characteristics controls* | yes | yes | yes | yes | yes | yes |
| *Estab. controls:* age, size, region | yes | yes | yes | yes | yes | yes |
| *Establishment controls:* Industries | 3-digit | 3-digit | 3-digit | 3-digit | 3-digit | 3-digit |
| Observations | 32484 | 32484 | 32484 | 32484 | 32484 | 32484 |
| R-squared | 0.062 | 0.233 | 0.070 | 0.118 | 0.163 | 0.218 |
| Mean of the dependant variable | 0.495 | 0.884 | 0.180 | 0.133 | 2.032 | 3.975 |

*Notes*. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0. Workers' characteristics controls are the percentage of blue collars, clerks, technicians and managers as well as the percentage of women and workers above 40 years old at the establishment level. Establishment controls are identical to those used in previous tables.

[32] Since the DMMO-EMMO files do not report these establishment characteristics, they are therefore drawn from the REPONSE survey and thus refer to 2004. The average workers' characteristics at the establishment level come from the social security records DADS (see data section) and have been matched with the REPONSE dataset by the Ministry of Labor.